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INTER-REGIONAL VOLATILITY SPILLOVERS BETWEEN EMERGING CAPITAL MARKETS: EVIDENCE FROM TURKEY AND BRAZIL

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Inter-Regional Volatility Spillovers between Emerging Capital Markets: Evidence from Turkey and Brazil*

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Abstract

This study investigates volatility spillovers between two stock markets, Turkish and Brazilian, located in different regions of the world. Using a misspecification robust causality-in-variance test, we find strong evidence supporting volatility spillovers from Istanbul Stock Exchange (ISE) to São Paulo Stock Exchange (BOVESPA). The results imply that financial crises may change the size and the direction of volatility spillovers between markets.

Keywords: causality-in-variance, volatility spillovers, emerging markets, Turkey, Brazil.

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1 Introduction

Volatility and return spillovers between emerging capital markets have been subject to extensive empirical research. Most of the related research focus on volatility spillovers in the context of financial crises. Obviously, volatility spillovers are closely related to the transmission of shock via newly opened channels associated with crisis events called "contagion."[1] Therefore, uncovering the nature of volatility spillovers contributes to the research on the contagion of financial crises as well. Understanding volatility spillovers is important for portfolio diversification and hedging strategies i.e. investor behavior. Bekaert and Harvey (2003) argue that greater integration of international stock markets and correlated stock price volatility decreases the opportunities for international portfolio diversification. Analyzing the transmission of volatility may also shed light on the nature of information flows between international markets. King and Wadhwani (1990) explain the volatility spillovers by the rational attempts of agents to use imperfect information about the events relevant to stock prices.

Most studies focus on volatility spillovers between developed stock markets and emerging stock markets, or between emerging markets located on the same region with strong real economic and financial linkages. Volatility spillovers among the emerging capital markets in the same region have found theoretical and empirical support, while the spillovers among distant emerging capital markets needs more work both theoretical and empirical. Theoretically, advances in information technologies, capital mobility, competition on product markets of third countries and similarities in asset structures might cause transmission of volatility between two capital markets located on different regions, even though they have no significant real and financial linkages. In the absence of strong trade and financial links, explanations of inter-regional volatility spillovers should rest on investor behavior and information flows. One such explanation is herd behavior. Calvo and Mendoza (2000), for instance, argue that in the presence of fixed information costs it might be rational for market participants to mimic other markets or investors that they think have more information. Another source of volatility spillovers between emerging markets might be the linkages with developed markets. According to Calvo (1999), developed stock markets can act as a conduit for volatility across emerg-

[1]We use the term "spillover" instead of "contagion" since it implies transmission of movements in general. It does not preclude the contagion (For more on the definition of contagion, see Forbes and Rigobon, 2002; Dungey et al, 2005).

ing markets in different regions. Along the same line, Kaminsky and Reinhart (2008) analyze the transmission of financial turmoil among emerging countries located in different regions through financial centers. In a recent study, Dungey and Martin (2007) provide empirical evidence for the role of developed markets in volatility transmission across emerging markets.

Empirical studies seem to support the above conjectures on volatility spillovers across regions. Fujii (2005) reports evidence in favor of volatility spillovers from Asian emerging markets to Latin American markets. Specifically, the author identifies volatility spillovers from Thailand to Argentina, Brazil, Chile and Mexico, and from Hong Kong and the Philippines to Mexico. Using a similar methodology, Gebka and Serwa (2007) find mixed results on volatility spillovers among the emerging capital markets in Eastern Europe, East Asia and Latin America.

This paper extends the literature on inter-regional volatility spillovers by providing empirical evidence from Turkish and Brazilian stock markets using the causality-in-variance test developed first by Cheung and Ng (1996), and further improved by Hong (2001). Using the conditional variances obtained from univariate Generalized Autoregressive Conditional Heteroskedasticity (GARCH) estimations, we investigate the existence and the direction of volatility spillovers between ISE (Istanbul Stock Exchange) and BOVESPA (São Paulo Stock Exchange) located in distant regions. Our findings mainly confirm the conclusions from other studies on volatility spillovers across regions. There exist volatility transmission between the Brazil and the Turkish markets, even though the two countries have no strong real and financial linkages. Moreover, financial crises seem to change the nature of this spillover effects.

Located in distance regions, Turkey and Brazil are not among the first 40 trading partners with each other. On the other hand, similarities between the two countries make them the subjects of empirical studies.[2] Both countries are considered emerging markets. Turkey and Brazil are both open economies with international capital flows allowed. Turkey has liberated its capital account in 1989 and Brazil has done the same in 1991 (Bekaert and Harvey, 2000). Turkey and Brazil have been heavily indebted to IMF, and both have similar economic experiences (Metin and Muradoglu, 2001). We observe ample anecdotal evidence from ISE participants pronouncing that they

[2]For some other studies on the two countries, see, for instance, Iwata and Tanner (2003), Celasun et al. (2003), Tanner and Samake (2006) and Baig and et al. (2000;2006).

closely monitor BOVESPA, suggesting the possibility of information flow between two markets[3]. In fact, the empirical research seems to suggests a significant relationship between the mean returns of the two stock markets. In two recent studies, Erbaykal et al. (2008) and Yalama (2009) find a long-run relationship between BOVESPA and ISE mean returns. They do not, however, investigate the causal nature of this relationship. Their findings naturally raises the question about possible volatility spillovers between the two markets.

Alper and Yilmaz (2004), which is the only study, to our knowledge, analyzing the possible volatility spillovers between ISE and BOVESPA, find no spillover effect between the volatilities of the two markets. Their study differs from the present work in terms of data and methodology. They use weekly return data and a parametric bivariate GARCH methodology, which is known to suffer from distributional and modeling misspecifications (Hafner and Herwartz, 2004). The present paper is closely related with two studies on inter-regional volatility spillovers, Fuji (2005) and Gebka and Serwa (2007). Both papers employ the causality-in-variance test proposed by Cheung and Ng (1996) in order to study volatility spillovers. Gebka and Serwa (2007) also find volatility spillovers among the emerging capital markets in Eastern Europe, East Asia and Latin America. These authors detect volatility spillovers from Latin American to European emerging markets, but not the otherwise, with the exception of Russian and Argentinean stock markets. One finding of Gebka and Serwa (2007) is that crises periods do not differ regarding volatility spillovers across markets. Neither of these studies includes the Turkish stock market.

This study reexamines the volatility spillovers between ISE and BOVESPA using daily data from 1993-2009 and the cross-correlation based testing methodology by proposed by Hong (2001). We investigate the volatility spillovers between two markets for the whole period, as well as the sub-periods divided by the major financial crises in Brazil. In our testing procedure, we try to account for the financial center effect as suggested by Kaminsky and Reinhart (2008). In fact, Ozun (2007) reports that

[3] The anecdotal evidence include interviews with brokers and market participants, and numerous internet sources and newspaper columns. See for example, "Brazilian bourse Bovespa and Istanbul stock index decouple," Turkish Daily News, Thursday, May 8, 2008, url: http://www.turkishdailynews.com.tr/article.php?enewsid=103913, and "Brezilya sarsıldı İMKB düştü, dolar 1.35 YTL'yi gördü (Brazil shaked, ISE down, USD see 1.35YTL)," Sabah Dailly Newspaper, July 27, 2005, url: http://arsiv.sabah.com.tr/2005/07/27/eko118.html.

developed markets effect the volatility of both BOVESPA and ISE. In contrast to Gebka and Serwa (2007), we find that pre-crisis and crisis periods differ in terms of spillover effects. Although it is not the main focus of the paper, we also examine the spillovers of volatility between the developed markets, namely the US and the UK markets, and the two emerging markets considered here.

2 Econometric Methodology

We employ the two-step causality-in-variance test suggested by Hong (2001) originally proposed by Cheung and Ng (1996). The test is argued to have high power and be robust to distributional assumptions. The test procedure is based on cross correlations of conditional variances obtained by a univariate GARCH process. This approach has the advantage of identifying the direction of volatility transmission in addition to detecting the existence of such spillovers. Unlike multivariate GARCH approaches to spillovers, the causality-in-variance test is not affected by specification errors. Hence, the results are robust to non-normal error terms. Moreover, the testing procedure is not subject to the generated regressors problem pointed out by Pagan (1984) since they are not regression based, unlike the Granger-causality test introduced by Granger (1969,1980).

The version of the test by Hong (2001) has two main advantages over the one originally proposed by Cheung and Ng (1996). First, it has more power in the case of large lag values. Second, Hong's (2001) version allows including the returns from the other market in the mean equation of GARCH model. This, in the author's words, "filters out possible effects of causality-in-mean."

The testing procedure involves estimating univariate GARCH models and applying the test statistic distributed asymptotically standard normal to the standardized conditional variances. As the first step in the testing procedure, we start with modeling the return series from both markets using the GARCH approach by Bollersev (1986). In order to filter out any causality-in-mean effect, we include the one-lagged returns from the other market. The following is the general form of the model employed in modeling both markets:

$$r_t^X = \alpha_0 + \sum_{i=1}^{m} \alpha_i r_{t-i}^X + \sum_{j=1}^{n} \theta_j r_{t-j}^Y + z_t \tag{1}$$

$$z_t = \varepsilon_t \sqrt{h_{it}}, \qquad \varepsilon_t \backsim iid(0,1)$$

$$h_t = \beta_0 + \sum_{i=1}^{p} \beta_i h_{t-i} + \sum_{j=1}^{q} \beta_j \varepsilon_{t-j}^2$$

where r_t^X is the return on the market being modeled, and r_t^Y is the return on the other market. The standardized disturbances are assumed to be independently, identically and normally distributed with zero mean and unit variance. We estimate the conditional variances and residuals using the quasi-maximum likelihood method.

In the second step, we construct the null hypothesis and the test statistic as follows. Let $\{R_{1t}, R_{2t}\}_{t=-\infty}^{\infty}$ be two stationary return series that we would like to test for causality in their time-varying conditional variances, and, I_{it}, $i = 1, 2$ be the information set defined as $I_{it} = \{R_{ij}, j \geq 0\}$, $I_t = I_{1t} \cup I_{2t}$

Assume the disturbance process is

$$z_{it} = \varepsilon_{it} \sqrt{h_{it}} \tag{2}$$

where h_{it} is a positive time-varying measurable function with respect to $I_{it-1,}$ and ε_{it} is an innovation process with $\mathrm{E}(\varepsilon_{it} \,|I_{it-1}) = 0$ and $\mathrm{E}(\varepsilon_{it}^2 \,|I_{it-1}) = 1$

Following Hong (2001), the null hypothesis that R_{2t} does not cause R_{1t} in variance can be written as:

$$H_0: \ \mathrm{Var}(z_{1t} \,|I_{1t-1}) = \mathrm{Var}(z_{1t} \,|I_{t-1}) \tag{3}$$

If H_0 is rejected, we say that R_{2t} causes R_{1t} in variance.

In order to construct the test statistic, define the centered squared standardized residuals from the GARCH (p,q) estimation as

$$\hat{\mu}_t = \hat{z}_{1t}^2 / \hat{h}_{1t} \qquad and \quad \hat{\nu}_t = \hat{z}_{2t}^2 / \hat{h}_{2t} \tag{4}$$

and the sample cross-correlation at lag k,

$$r_{uv}(k) = c_{uv}(k) \ / \sqrt{c_{uu}(0).c_{vv}(0)}, \tag{5}$$

where the sample cross covariances

$$c_{uv}(k) = \begin{cases} T^{-1}, \sum_{t=k+1}^{T-1} \hat{u}_t \hat{v}_{t-k}, & k \geq 0 \\ T^{-1}, \sum_{t=-k+1}^{T-1} \hat{u}_{t+j} \hat{v}_t, & k < 0 \end{cases} \tag{6}$$

with T being the sample size$c_{uu}(0) = T^{-1}, \sum_{t=1}^{T} \hat{u}_t^2$ and $c_{vv}(0) = T^{-1}, \sum_{t=1}^{T} \hat{v}_t^2$.

Hong (2001) suggests the following test statistic:

$$Q = T \left\{ \sum_{k=1}^{T} w^2(k/M) r_{uv}^2(k) - c(w) \right\} / \{2D(w)\}^{1/2}, \tag{7}$$

$$C(w) = \sum_{k=1}^{T-1} (1 - k/T) w^2(k/M),$$

$$D(w) = \sum_{k=1}^{T-1} (1 - k/T)[1 - (j+1)/T] w^4(k/M)$$

where where M is the number of cross correlations included. We can think of M as the lags considered for the spillover effect. The function w(.) is a weighting function for which we used the Bartlett, Daniell and Truncated kernels:

Bartlett:

$$w(k/M) = \begin{cases} 1 - |k/M|, & |k/M| \leq 2 \\ 0, & \text{otherwise} \end{cases}$$

Truncated:

$$w(k/M) = \begin{cases} 1, & |k/M| \leq 2 \\ 0, & \text{otherwise} \end{cases}$$

Daniell:

$$w(k/M) = \sin(\pi k/M)/(\pi k/M), -\infty < (k/M) \leq \infty$$

Note that the test statistic uses the complete "bivariate" information set $I_{t-1,}$ so that any causality-in-mean is filtered out. We achieve this by including the returns from the other market in the mean equation of the GARCH model. In addition, in order to take the effects of developed markets such as the US and the UK markets into account, we also included the return series of these markets in the mean equation of (1). Hong (2001) shows that Q is a one-sided test statistic, and distributed, under the null, asymptotically as standard normal. The null hypothesis given by (3) can be tested by calculating the test statistic Q and comparing it with the upper tail probabilities of N(0,1).

3 Data and Empirical Findings

The data set covers the period from April 09, 1993 to April 10, 2009. We obtained all the series from DataStream. Definitions of the series are in Table 1. Following Diebold and Yilmaz (2009), among others, we use local currency dominated return series, since we are interested in volatility faced by domestic participants.[4] In all analyses in the study, we used return series calculated by taking logarithmic differences and by multiplying by 100. As a preliminary analysis, we check the stationarity of index returns by ADF (Augmented Dickey-Fuller) unit root test, and all return series we found to be stationary. Unit-root test results are not reported but are available from the authors upon request.

In the first step of testing for causality-in-variance between the markets, we need to perform the GARCH estimations. For brevity we omit the estimated GARCH results, but they are available on request; most are GARCH(1,1) or GARCH(2,2) models selected using AIC criterion. We aim to test volatility spillovers between the Brazilian and the Turkish markets, as well as the spillovers between Brazil-US, Brazil-UK, Turkey-US and Turkey-UK. Hence we estimate six AR-GARCH models. All models include the US and the UK return series. The reason we include the US and the UK markets is to eliminate any indirect volatility spillovers via these financial centers as pointed out by Kaminsky and Reinhart (2008). Among three financial centers we originally considered, namely the US, Europe and Japan, we choose to include only the US and European (UK) markets, since most of the lending from Japan mainly goes to Asian countries (van Rijckeghem and Weder, 2003). Specifically, for Brazil-Turkey case, the Brazil model includes UK, US and Turkey with one lag and AR terms in the

[4] By using returns dominated in local currency, we implicitly assume that international and domestic market participants are able to, at least partially, hedge their foreing exchange risks. In our opinion, assuming they are not able to hedge their exchange rate risks at all, by using returns denominated in a common currency would introduce greater bias into the analysis.

mean equation. Similarly, the Turkey model includes UK, US and Brazil with one lag and AR terms. For developed-emerging pairs, developed market models, US and UK, include AR terms, the other developed market with one lag and the emerging market with one lag. For example, the US model, for the US-Brazil pair, has UK and Brazil with one lags and AR terms.

In the second step, we compute the one sided test statistic (7) for the null hypotheses: (i) "MARKET 1 does not cause MARKET 2 in variance" (ii) "MARKET 2 does not cause MARKET 1 in variance." The rejection of the null (i) means that there exist volatility spillovers from MARKET 1 to MARKET 2. Similarly, the rejection of the null (ii) means that there exist volatility spillovers from MARKET 2 to MARKET 1.

We present the test results for both hypotheses using Barlett, Truncated and Daniell kernels in Table 2-6. In both cases, we choose a maximum of $M = 15$. Considering the fact that we can interpret M as number of lags in cross correlations, 15 is sufficiently long for daily observations. In fact, we observed that the test results with $M > 15$ do not change significantly. The tables report only the upper tail probabilities for standard normal distribution; rejection probabilities under 0.05 are in bold-face. Table 2 reports the test statistic and corresponding p-values for the null hypotheses for Brazil and Turkey. The results suggest bidirectional causality in variance with the spillover from Brazil to Turkey more pronounced. The rejection of the null that Turkey does not cause Brazil seems weaker. These findings differ from Alper and Yilmaz (2004) who do not find any spillover of volatility between these two markets.

Table 3 and 4 reports the results for Brazil-UK and Brazil-US respectively. In these cases, again, we can infer two-way causality. The test results, however, seem to sensitive the choice of the kernel functions. The causality from Brazil to US, on the other hand, is robust to the kernel function choice. This observation may suggest that the Brazilian market exports volatility to the US market.

While the Brazilian market interacts with the two developed markets in both directions, the Turkish market seems to import volatility from the developed markets. Table 5 and 6 shows that there is no significant causality in variance from Turkey to neither the US nor the UK. The US and the UK markets export significant volatility spillovers to Turkey. This result is not surprising considering the size and the influence of these markets, and justifies our inclusion of the two financial centers in GARCH estimations.

In order to see whether financial crises have any effect on the nature of the volatility spillovers between the two emerging markets, we divide the sample period into two sub-periods according to a particular financial crisis. Our choice of crisis is the 1999 Brazilian financial meltdown. Following Dungey et al. (2005; 2010), we chose the start of the Brazilian crisis as January 7, 1999, before the effective devaluation of the Real on January 15, 1999. The crisis was triggered by the announcement of a 90-day moratorium on debt payments to the central government by a provincial governor on January 6, 1999. The announcement raised the worries of investors causing a rapid capital outflow. The events, eventually, lead to the devaluation of real by the Central Bank of Brazil.[5] Table 7 presents some descriptive statistics of the crisis, pre-crisis, and total periods for the four countries. It is easily seen in Table 7 that the mean returns of equity market for all countries decrease between the stable and crisis periods. Table 8 reports the covariances between the Brazilian and the Turkish markets in pre-crisis and the crisis periods. There is a large increase in the covariances in the crisis period which indicates an interdependence from contagion. This is true also for other markets.

For the crisis period we construct a crisis dummy taking the value of 1 during the exogenously defined crisis period and 0 otherwise, and re-estimate the AR-GARCH models for both sub-periods. We show the pre-crisis period volatility spillovers in Table 9. Although the spillover from the Brazilian market is more robust to the kernel function choice, we can safely conclude the two way causality-in-variance between the markets. This conclusion, however, is not true for the crisis period. In the crisis period, we observe volatility spillovers only from Brazil to Turkey (see Table 10) [6]. These results are in contrast with the findings of Gebka and Serwa (2007). They find no difference between the sub-periods of crisis regarding the inter-regional volatility spillovers among the countries analyzed.

[5] See Ferreira and Tullio (2002) for more on the effects of this specific currency crisis on Brazilian economy.

[6] For further evidence, see Baig and Goldfajn, 1999; Forbes and Rigobon, 2002; Dungey et al, 2005. They show that correlations in markets increase significantly during the crisis period.

4 Conclusion

The existence of volatility spillovers between emerging stock markets located in different regions with no sizable real economic and financial linkages has implications regarding international flows of information. Considering a special case, this study looks for evidence on volatility spillovers between two emerging markets, BOVESPA and ISE, located in different regions and with insubstantial trade and financial interaction. By employing a cross-correlation based causality-in-variance test, we test for the existence and the direction of volatility transmissions between the two countries in stable as well as the crisis periods. Our model allows us to control the developed country or financial center effects. Therefore, we can interpret our findings as direct linkages between two stock markets.

The findings indicate the transmission of volatility between BOVESPA and ISE in both directions. Moreover, causality of the volatility runs only one way in the crisis period: from BOVESPA to ISE. This result suggests that large shocks in São Paulo Stock Exchange increase the volatility in Istanbul Stock Exchange especially in the periods of financial crises. This is important in the sense that the nature of the relationship between the two emerging markets changes significantly in crises periods.

We can confidently discard the explanations based on trade links between Turkey and Brazil. Another explanation we can put less weight is the financial center effects, since we control for financial centers in our analysis. One reasonable explanation of the findings might be the financial links between two countries. These financial links can be in the form of international investors or common lenders such as banks. van Rijckeghem and Weder (2003) report that European banks have lending on Latin America and Eastern Europe and Asia in almost equal proportions, while the US banks concentrated mainly on Latin America especially after the Asian crisis. Another reasonable conjecture might be based on information flows. It is not unreasonable to think about domestic investors in both countries following the other market's movements closely due to costly information. Gathering and processing international information is costly, and the cheapest way to make use of this information is to follow the markets that resembles each other in many ways. Many domestic players, for instance, in İstanbul Stock Exchange explicitly pronounce that they closely monitor the ISE data on international participants. Assessing the relative weights of the above explanations calls for more research on the issue.

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Table 1:
Data Definition: All data are from DataStream

Country	Name of Index	Series
TURKEY	ISE NATIONAL 100 PRICE INDEX	TRKISTB
BRAZIL	BRAZIL BOVESPA PRICE INDEX	BRBOVES
UK	FTSE 100 PRICE INDEX	FTSE100
US	S&P 500 COMPOSITE PRICE INDEX	S&PCOMP

Table 2:
Results of Hong test for total period: Brazil and Turkey

	TURKEY$\nrightarrow$BRAZIL			BRAZIL$\nrightarrow$TURKEY		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.620	**0.020**	-	0.401	0.344
2	0.620	0.758	0.385	0.401	**0.000**	0.404
3	0.678	**0.005**	0.692	**0.000**	**0.000**	**0.000**
4	0.501	**0.030**	0.501	**0.000**	**0.000**	**0.000**
5	0.296	0.083	0.241	**0.000**	**0.000**	**0.000**
6	0.194	0.063	0.115	**0.000**	**0.000**	**0.000**
7	0.143	0.077	0.082	**0.000**	**0.000**	**0.000**
8	0.115	**0.034**	0.069	**0.000**	**0.000**	**0.000**
9	0.096	**0.018**	0.060	**0.000**	**0.000**	**0.000**
10	0.080	**0.009**	0.059	**0.000**	**0.000**	**0.000**
11	0.066	**0.012**	**0.049**	**0.000**	**0.000**	**0.000**
12	0.055	**0.020**	**0.038**	**0.000**	**0.000**	**0.000**
13	**0.046**	**0.018**	**0.031**	**0.000**	**0.000**	**0.000**
14	**0.039**	**0.032**	**0.026**	**0.000**	**0.000**	**0.000**
15	**0.035**	0.054	**0.022**	**0.000**	**0.000**	**0.000**

Notes :The values represents upper tail probabilities of standart normal distribution

Table 3:
Results of Hong test for total period: Brazil and UK

	UK↛BRAZIL			BRAZIL↛UK		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.707	0.294	-	0.691	0.605
2	0.707	0.623	0.725	0.691	**0.011**	0.649
3	0.693	0.228	0.254	0.334	0.063	0.639
4	0.610	**0.017**	0.135	0.157	**0.000**	0.614
5	0.456	**0.045**	**0.051**	0.071	**0.003**	0.434
6	0.317	**0.045**	**0.017**	**0.031**	**0.009**	0.280
7	0.227	0.066	**0.009**	**0.017**	**0.024**	0.173
8	0.172	0.100	**0.006**	**0.012**	**0.013**	0.123
9	0.140	0.158	**0.005**	**0.010**	**0.025**	0.092
10	0.122	0.153	**0.005**	**0.008**	**0.028**	0.080
11	0.112	0.214	**0.005**	**0.008**	**0.021**	0.077
12	0.106	0.200	**0.005**	**0.007**	**0.037**	0.073
13	0.104	0.195	**0.005**	**0.007**	0.062	0.075
14	0.104	0.251	**0.007**	**0.007**	0.095	0.084
15	0.104	0.276	**0.008**	**0.008**	0.067	0.094

Notes :The values represents upper tail probabilities of standart normal distribution

Table 4:
Results of Hong test for total period: Brazil and US

	US↛BRAZIL			BRAZIL↛US		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.670	0.926	-	0.741	0.132
2	0.670	**0.008**	0.628	0.741	**0.000**	0.711
3	0.307	**0.010**	0.296	**0.000**	**0.000**	**0.000**
4	0.112	**0.019**	0.083	**0.000**	**0.000**	**0.000**
5	0.055	**0.008**	**0.037**	**0.000**	**0.000**	**0.000**
6	**0.032**	**0.025**	**0.019**	**0.000**	**0.000**	**0.000**
7	**0.023**	0.051	**0.012**	**0.000**	**0.000**	**0.000**
8	**0.019**	0.093	**0.011**	**0.000**	**0.000**	**0.000**
9	**0.017**	0.067	**0.013**	**0.000**	**0.000**	**0.000**
10	**0.018**	0.051	**0.015**	**0.000**	**0.000**	**0.000**
11	**0.018**	0.067	**0.017**	**0.000**	**0.000**	**0.000**
12	**0.019**	**0.044**	**0.018**	**0.000**	**0.000**	**0.000**
13	**0.019**	0.055	**0.021**	**0.000**	**0.000**	**0.000**
14	**0.020**	0.086	**0.023**	**0.000**	**0.000**	**0.000**
15	**0.020**	0.121	**0.024**	**0.000**	**0.000**	**0.000**

Notes :The values represents upper tail probabilities of standart normal distribution

Table 5:
Results of Hong test for total period: Turkey and UK

	UK↛TURKEY			TURKEY↛UK		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.725	0.868	-	0.252	**0.002**
2	0.725	**0.000**	0.089	0.252	0.510	0.275
3	0.137	**0.000**	0.118	0.316	0.615	0.333
4	**0.000**	**0.000**	**0.000**	0.391	0.652	0.359
5	**0.000**	**0.000**	**0.000**	0.453	0.695	0.479
6	**0.000**	**0.000**	**0.000**	0.503	0.247	0.536
7	**0.000**	**0.000**	**0.000**	0.518	0.348	0.557
8	**0.000**	**0.000**	**0.000**	0.508	0.372	0.553
9	**0.000**	**0.000**	**0.000**	0.494	0.448	0.512
10	**0.000**	**0.000**	**0.000**	0.482	0.373	0.489
11	**0.000**	**0.000**	**0.000**	0.473	0.341	0.475
12	**0.000**	**0.000**	**0.000**	0.462	0.423	0.466
13	**0.000**	**0.000**	**0.000**	0.454	0.433	0.457
14	**0.000**	**0.000**	**0.000**	0.447	0.504	0.447
15	**0.000**	**0.000**	**0.000**	0.442	0.230	0.434

Notes :The values represents upper tail probabilities of standart normal distribution

Table 6:
Results of Hong test for total period: Turkey and US

	US↛TURKEY			TURKEY↛US		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.591	0.887	-	0.531	0.772
2	0.591	**0.000**	0.618	0.531	0.108	0.566
3	0.132	**0.004**	0.133	0356	0.271	0.353
4	**0.025**	**0.010**	**0.015**	0.275	0.152	0.202
5	**0.010**	**0.016**	**0.006**	0.237	0.242	0.203
6	**0.007**	**0.042**	**0.003**	0.216	0.017	0.191
7	**0.006**	0.045	**0.004**	0.186	0.038	0.177
8	**0.006**	0.086	**0.005**	0.152	0.072	0.144
9	**0.007**	0.111	**0.007**	0.127	0.033	0.116
10	**0.008**	0.055	**0.009**	0.108	0.044	0.095
11	**0.009**	0.088	**0.012**	0.093	0.067	0.083
12	**0.011**	0.122	**0.014**	0.082	0.106	0.075
13	**0.012**	0.177	**0.018**	0.075	0.059	0.069
14	**0.014**	0.197	**0.023**	0.070	0.086	0.064
15	**0.016**	0.185	**0.027**	0.066	0.064	0.062

Notes :The values represents upper tail probabilities of standart normal distribution

Table 7:
Descriptive statistics of daily percentage equity returns for selected periods: Pre-crisis period (09th April 1993 to 06th January 1999), Crisis period (7th January 1999 to 10th April 2009), Total period (7th January 1999 to 10th April 2009).

Country	Sample period	Mean	Mak	Min	Sdev	Skewness	Kurtosis
TURKEY	Pre-crisis period	0.251	15.648	-16.167	3.057	-0.303	5.658
	Crisis period	0.086	17.774	-19.979	2.652	0.107	8.235
	Total period	0.145	17.774	-19.979	2.805	-0.071	7.061
BRAZIL	Pre-crisis period	0.399	22.813	-17.229	3.192	0.141	7.505
	Crisis period	0.068	28.818	-12.096	2.104	0.829	18.979
	Total period	0.187	28.818	-17.229	2.553	0.476	12.166
UK	Pre-crisis period	0.052	4.345	-3.661	0.865	-0.087	5.399
	Crisis period	-0.016	9.384	-9.266	1.316	-0.128	9.415
	Total period	-0.008	9.384	-9.266	1.175	-0.157	10.092
US	Pre-crisis period	0.071	4.989	-7.133	0.855	-0.647	12.403
	Crisis period	-0.015	10.957	-9.470	1.362	-0.082	11.129
	Total period	0.016	10.957	-9.470	1.205	-0.196	12.751

Table 8:
Covariance of daily percentage equity returns for selected periods: Pre-crisis period (09th April 1993 to 06th January 1999), Crisis period (7th January 1999 to 10th April 2009), Total period (7th January 1999 to 10th April 2009).

Country	TURKEY	BRAZIL	UK	US
	Pre-crisis period			
TURKEY	9.337			
BRAZIL	0.707	10.180		
UK	0.196	0.984	0.747	
US	0.463	0.610	0.278	0.730
	Crisis period			
TURKEY	7.031			
BRAZIL	1.258	4.427		
UK	0.564	1.630	1.732	
US	1.012	1.179	0.862	1.854
	Total period			
TURKEY	7.865			
BRAZIL	1.073	6.516		
UK	0.435	1.405	1.380	
US	0.818	0.980	0.654	1.452

Table 9:
Results of Hong test for Pre-crisis period: Brazil and Turkey

	TURKEY$\nrightarrow$BRAZIL			BRAZIL$\nrightarrow$TURKEY		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	**0.031**	0.576	-	0.601	0.051
2	**0.031**	0.170	**0.019**	0.601	**0.000**	0.552
3	**0.046**	**0.035**	**0.044**	**0.007**	**0.000**	**0.005**
4	**0.048**	0.096	0.053	**0.000**	**0.000**	**0.000**
5	**0.045**	0.168	**0.036**	**0.000**	**0.000**	**0.000**
6	**0.049**	0.260	**0.041**	**0.000**	**0.000**	**0.000**
7	0.056	0.075	0.060	**0.000**	**0.000**	**0.000**
8	0.063	**0.015**	0.073	**0.000**	**0.000**	**0.000**
9	0.062	**0.014**	0.077	**0.000**	**0.000**	**0.000**
10	0.057	**0.011**	0.076	**0.000**	**0.000**	**0.000**
11	**0.050**	**0.024**	0.063	**0.000**	**0.000**	**0.000**
12	**0.044**	**0.030**	**0.051**	**0.000**	**0.000**	**0.000**
13	**0.039**	**0.019**	**0.047**	**0.000**	**0.000**	**0.000**
14	**0.035**	**0.026**	**0.041**	**0.000**	**0.000**	**0.000**
15	**0.032**	**0.042**	**0.036**	**0.000**	**0.000**	**0.000**

Notes :The values represents upper tail probabilities of standart normal distribution

Table 10:
Results of Hong test for crisis period: Brazil and Turkey

	TURKEY$\nrightarrow$BRAZIL			BRAZIL$\nrightarrow$TURKEY		
M	Bartlett	Truncated	Daniell	Bartlett	Truncated	Daniell
1	-	0.737	0.260	-	0.493	0.944
2	0.737	0.779	0.636	0.493	**0.000**	0.473
3	0.766	0.152	0.794	**0.025**	**0.000**	**0.018**
4	0.681	0.288	0.692	**0.000**	**0.000**	**0.000**
5	0.572	0.405	0.566	**0.000**	**0.000**	**0.000**
6	0.507	0.328	0.473	**0.000**	**0.000**	**0.000**
7	0.467	0.430	0.421	**0.000**	**0.000**	**0.000**
8	0.442	0.523	0.404	**0.000**	**0.000**	**0.000**
9	0.430	0.468	0.399	**0.000**	**0.000**	**0.000**
10	0.427	0.479	0.415	**0.000**	**0.000**	**0.000**
11	0.426	0.464	0.415	**0.000**	**0.000**	**0.000**
12	0.426	0.546	0.411	**0.000**	**0.001**	**0.000**
13	0.429	0.576	0.418	**0.000**	**0.002**	**0.000**
14	0.433	0.636	0.431	**0.000**	**0.002**	**0.000**
15	0.439	0.695	0.441	**0.000**	**0.004**	**0.000**

Notes :The values represents upper tail probabilities of standart normal distribution